Transitions in welfare participation and female headship

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Abstract:
This study uses data from the 1990, 1992, 1993 and 1996 panels of the Survey of Income and Program Participation to examine how welfare policies and local economic conditions contribute to women's transitions into and out of female headship and into and out of welfare participation. It also examines whether welfare participation is directly associated with longer spells of headship. The study employs a simultaneous hazards approach that accounts for unobserved heterogeneity in all of its transition models and for the endogeneity of welfare participation in its headship model. The estimation results indicate that welfare participation significantly reduces the chances of leaving female headship. The estimates also reveal that more generous welfare benefits do not directly contribute to headship but rather contribute indirectly to headship by increasing the chances that a mother will enter welfare and consequently remain a single mother for longer. More generous Earned Income Tax Credit benefits are associated with more stable arrangements for both headship and welfare participation. Other measures of welfare policies, including indicators for the adoption of welfare waivers and the implementation of Temporary Assistance for Needy Families programs, are generally not significantly associated with headship or welfare receipt. Better economic opportunities are estimated to increase headship but reduce welfare participation among unmarried mothers.
Keywords: Female headship, Hazard models, Welfare participation

Article:
Introduction
Policy makers have long expressed concern over the linkage between welfare use and family structure. In the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996, Congress and the President sought to change public assistance policies so that they would not only promote economic self-sufficiency but also support marriage and discourage single parenthood. In the current proposals to reauthorize PRWORA, family structure has become even more central.

Researchers, too, have been keenly interested in understanding how welfare policies affect demographic outcomes including births, marriages and living arrangements. The studies in this area have generally examined reduced-form associations between policies and demographic behaviors. Few of the studies on demographic outcomes have considered how policies might operate through welfare participation or how participation might directly affect demographic behavior. The chief difficulty in such an analysis is that welfare participation and family structure are both endogenous variables. An analysis of the direct relationship between welfare participation and family structure must account for unobserved factors that might contribute to both outcomes. Reduced-form strategies sidestep this thorny issue.

In this article, we estimate transition models into and out of female headship and into and out of welfare participation that allow unobserved heterogeneity to be correlated among the various transitions. We further allow the transition out of female headship to directly depend on welfare participation status, thereby estimating a structural direct effect. We thus address whether being on welfare directly reduces the likelihood that a women leaves headship (which would be primarily by marriage) after properly controlling for individual heterogeneity. We employ Lillard's (1993) simultaneous hazard approach to estimate the model using individual-level data from the 1990, 1992, 1993 and 1996 panels of the Survey of Income and Program Participation (SIPP). These data are augmented by contextual data on state-level welfare policies, state and Federal Earned Income Tax Credit (EITC) policies, and county-level labor and marriage market conditions. Our empirical analysis reveals that participating in welfare reduces the chance of leaving female headship and that welfare policies affect headship through program participation.

Background and significance

Concern among policymakers and the public about rising rates of female headship in the US is not without foundation. Female-headed families tend to have higher poverty rates and more welfare usage than two- parent families (Lerman 1996). Children raised in female-headed families typically have worse schooling/developmental outcomes than those raised in two-parent families (Haveman & Wolfe 1994; McLanahan & Sandefur 1994).[1] Welfare reform legislation and its emphasis on encouraging two-parent families is a manifestation of this concern.

A large literature has developed concerning the impact of welfare programs on demographic decisions that give rise to female headship. In an influential book, Murray (1984) argued that public assistance programs were responsible for the growth in female headship. Murray pointed out that single parents were categorically eligible for AFDC, whereas married parents were generally ineligible. By subsidizing one family arrangement but not others, AFDC provided unambiguous incentives to become and remain a single parent. Subsequent changes in policies have reduced these incentives but have not eliminated them. The Family Support Act of 1998 required all states to offer AFDC-Unemployed Parent programs for married-couple families. These programs extended eligibility to two-parent families but required the primary earner to work less than 100 hours per month and have a recent history of work. The PRWORA allowed states to relax these requirements further; the TANF programs in 33 states now make no distinction between single-parent and married-couple families and base eligibility solely on financial considerations (Gardiner et al. 2002).

While the incentive effects of the old AFDC program seemed clear, there was considerable disagreement regarding the size of these effects. Time series evidence undercuts Murray's thesis regarding the centrality of public assistance policies. When adjusted for inflation, cash assistance has become less, not more, generous since the mid-1970s. If anything, this should have contributed to a decline, rather than a rise in headship.

Subsequent empirical research on the linkage between welfare benefits and family structure has been equivocal about the magnitude of effects. Moffitt (1998) reviewed a large number of studies on the impact of welfare benefit levels on fertility and marriage and concluded that welfare encouraged fertility and discouraged marriage, but that the sizes of the effects were likely small. Others have also surveyed the literature on marriage, cohabitation, fertility and divorce and reached similar conclusions (Acs 1995; Hoynes 1997b; Moffitt 1995, 2001; Ribar 1998).

Researchers have analyzed both experimental and observational evidence to investigate the impact of welfare programs on demographic outcomes. Most of the experimental evidence comes from states that conducted random-assignment evaluations of changes in welfare rules granted under program waivers. Although these evaluations were not primarily designed to investigate demographic impacts, they can provide very powerful evidence because they address problems of selectivity and policy endogeneity. Analyses of the Minnesota Family Investment Program by Knox et al. (2000) and Gennetian & Knox (2004) indicated that the program increased marriage and reduced divorce among participants. However, these appear to be isolated results. A meta analysis by Gennetian & Knox (2003) of waiver experiments in different states showed little consistent evidence of demographic effects. An analysis of a Canadian experiment by Harknett & Gennetian (2003) also produced inconclusive findings.

Observational studies have been able to examine the relationship between welfare policies and demographic outcomes using more general populations and across more varied environments. Schoeni & Blank (2000), Horvath-Rose & Peters (2001) and Bitler et al. (2004) have investigated the impacts of welfare using aggregate or state-level data and found effects of policies. A drawback to using aggregate data is that they suffer from composition effects whereby it is difficult to properly condition on individual traits. Aggregate studies also cannot control for duration effects.

A host of studies, including Blank (1999), Blank & Ruggles (1996), Fitzgerald (1995), Gittleman (2001), Klerman & Haider (forthcoming), and Ribar (2004), have used individual-level data to estimate models of welfare transitions. In their surveys of the literature, Blank (2002) and Moffitt (2002) have reported that most studies show that welfare benefits and labor market conditions have an impact on time spent on welfare. But the studies do not jointly consider female headship. Studies that consider family structure transitions such as Moffitt & Rendell (1995), Bitler et al. (2004), and Fitzgerald and Ribar (2004) do not jointly model welfare participation. In this paper we jointly model the two decisions.

Discrete choice models of demographic decisions and welfare have been used to jointly model the marriage/fertility and welfare choice. Duncan & Hoffman (1990) found little effect of benefit levels on births for black teens, and Hoffman & Duncan (1995) found little effect of benefits on divorce. Rosensweig (1999) found that higher AFDC benefits substantially increased the probability of a non-martial birth for low income women. His model allowed choices among three states – unmarried and childless, unmarried and with children, and married – and allowed for unobserved correlations in the utility in each state. He did not model the direct effect of participation in welfare.

Keane & Wolpin (2002) estimated a structural dynamic lifetime model that included welfare participation, fertility, marriage, work and school attendance using data from the NLSY79. They reported that welfare benefits had significant impacts on welfare participation, work and schooling decisions, but no significant effect on fertility and marriage decisions.

More closely related to our work, Teitler et al. (2003) have recently undertaken a preliminary analysis of the direct relationship between welfare participation and spells of unmarried motherhood. Their analysis followed mothers from the Fragile Families and Child Wellbeing study who were unmarried at the time of their children’s births and estimated hazard models of the women’s transitions into marriage. The models distinguished between women who were imputed to be eligible or ineligible for TANF as well as those who reported participating or not participating in the program. Their preliminary results, which did not account for the endogeneity of eligibility and participation, indicated that welfare participation was not strongly associated with marriage.

The endogeneity of the welfare participation decision could occur because unobserved characteristics of women prone to participate in welfare may also make them less likely to marry. Without proper controls for this heterogeneity, the unobserved characteristics could induce a spurious correlation between welfare use and headship transitions.

Our article extends the literature by using individual-level longitudinal data in a joint model of welfare and headship transitions. We allow these transitions to be linked by unobserved heterogeneity in a simultaneous hazard model. Furthermore, we allow headship transitions to depend directly on welfare participation. We estimate the impacts of welfare benefits, welfare waiver adoption and TANF adoption, and the EITC. Moffitt (1995, 1998) and others have noted that studies of welfare effects on demographic changes must be careful to control for the economic and policy environment across states so that welfare impacts do not become confounded with other changes. In this article we control for skill-specific county-level measures of the labor market and county- level marriage market variables to address these concerns.

Conceptual model

The empirical and conceptual analyses distinguish between women who are and are not single heads of families and women who do and do not participate in welfare. While these simplifications make the models much easier to work with, they also abstract from some relevant detail. Consider the routes into and out of headship. Women become female heads of families by bearing children out of wedlock or by dissolving marriages that have produced or include children. Mothers leave headship by marrying, having their children grow up, or having their children move out of the household. The relevant component behaviors regarding fertility, marriage and household structure can themselves be further broken down. For instance, a birth results from a series of outcomes involving sexual activity, contraception, completing a pregnancy, and keeping the child. The decision to marry follows some type of search activity and requires a corresponding decision from a partner.

Welfare participation also involves a complex set of processes. Families must apply for benefits and then be determined to be eligible. Eligible families who choose to take up benefits must further decide whether and how much to work, how to comply with program rules, and which services to receive. The potential interactions between the component processes of female headship and welfare participation are innumerable. Our conceptual analysis describes some of the reasons why such interactions arise, but it is far from exhaustive. More comprehensive theoretical discussions can be found in the articles by Blank & Ruggles (1996), Hoynes (1997a, b), Matthews et al. (1997), Peters et al. (2003) & Gennetian & Knox (2003).

Following Becker (1981), we examine female headship in a rational- choice framework. We also consider welfare participation in the same framework. The key assumptions of the rational-choice approach are that people evaluate how alternative decisions affect their well-being and choose actions that maximize their perceived well-being. Thus, a woman would become a single parent at a point in time if that outcome provides more expected lifetime utility than other family arrangements, including marriage or remaining childless. In making her decision, a woman would compare the immediate costs and rewards of single parenthood with the costs and rewards of the alternatives. She would also evaluate how the decision squares with her long-term interests. Decisions regarding welfare receipt would follow a similar calculus; mothers would participate in welfare if they perceive that it is their best interest to do so.

Without question, welfare participation affects the resources available to single mothers. The old AFDC program provided benefits to single mothers but not to married mothers or childless women. Thus, it created immediate incentives to become a single mother. Over the longer term, the incentives were less clear. On the one hand, the program could be viewed as subsidizing marital search and allowing women to be choosier regarding potential spouses. This would have also contributed to higher rates of headship. On the other hand, more selective searching could lead to better matches and more stable marriages. Or the availability of welfare could reduce the risk associated with a bad match. In addition, welfare participation itself might reduce marriage because, when looking ahead to marriage, the participants forgo the expected future benefits but non-participants forgo these same benefits net of the fixed cost and stigma cost of establishing benefits. Hence, marriage has a higher opportunity cost for current welfare participants.

The AFDC-UP and TANF programs, which are still conditioned on the presence of children, continue to provide incentives to become a parent. However, because they provide resources to both married and unmarried parents, the incentives for single parenthood are less clear. To the extent that the programs place more hurdles in front of married- couple families than single-parent families, they would create incentives to become a single parent. The incentives, however, would be milder than under the old AFDC program. Even if explicit hurdles are not present, it might be more difficult for married couples to qualify financially. This would be especially true of dual-earner couples.

By a similar logic, changes in welfare benefits or program rules would affect the value of participating in the program and alter the resources available to different types of families. Thus, changes in welfare policy variables, such as reductions in benefits, the imposition of family caps or time limits, or requirements to work, could affect female headship directly through welfare participation and indirectly through effects on potential future receipt.

The rational-choice framework also indicates that other variables will affect headship and welfare participation decisions. Labor market conditions would seem to be especially relevant. A woman living in an area with good job opportunities can expect higher earnings in both the married and unmarried states. Better earnings opportunities for women might reduce marriage through an "independence effect." Such an effect would arise if married women keep only a portion of an earnings increase while single women get to keep all of an increase. An independence effect would also occur if higher earnings subsidized marital search by single women. Higher earnings and employment could also change the probabilities of marriage by increasing women's bargaining power within marriage, increasing their exposure to employed men, or simply making them more financially attractive as potential spouses. Complicating this, better economic opportunities would also likely translate into higher male wages, which would raise the quality of marriage prospects and the possibility of marrying. Furthermore, higher wages and employment are likely to affect fertility by increasing the opportunity cost of having children but also increasing the resources available to raise them. Thus, higher wages and better employment prospects will have ambiguous effects on female headship; empirical research is needed to sort out the impacts.

The EITC also has ambiguous effects on marriage, as noted by Dickert-Conlin and Houser (1999). For single mothers with low earnings, the EITC can provide extra incentive to marry a man with earnings. For mothers with more earnings who already receive the EITC, the added income of a spouse could reduce or eliminate the EITC payment.

The effects of economic opportunities on welfare participation seem more clear cut because eligibility and benefit levels are both conditioned on incomes. What is less certain is whether the reforms initiated under waivers and then TANF policies strengthened or weakened this relationship. Higher earnings disregards and reduced benefit reduction rates in some states have made work and welfare more compatible. Indeed, the U.S. Department of Health and Human Services (2003) reported that almost 60% of TANF families in 2000 included at least one worker. At the same time, however, time limits may have made mothers more reluctant to participate in welfare and receive reduced benefits during periods when work is available. Women in these circumstances may prefer to "bank" their benefits as insurance against leaner times (Grogger 2003). While we expect better economic conditions to reduce welfare participation, the magnitude of the relationship is an open empirical question.

Data construction

Data preparation is divided into three tasks. We first use individual- level data on women from the SIPP to construct spells of headship and non-headship and spells of welfare participation and non-participation. These data also provide information on other personal and background characteristics of the women. Second, we augment this information with data on welfare policies and EITC benefits based on each woman's state of residence. Third, we add contextual variables about labor and marriage market conditions in the woman's county of residence. We restrict our sample to women aged 15-55.

Individual data from SIPP

We pool data from the 1990, 1992, 1993, and 1996 panels of SIPP. These data span the calendar period October 1989 to February 2000. This is an opportune period in which to observe behavioral responses to policy. During the 1990s states modified their welfare programs by obtaining waivers from the federal rules governing their programs. Many of these changes were incorporated into the 1996 PRWORA, though this bill also affected states that had not adopted waivers. In addition to the dramatic changes in welfare policies, the EITC was also adjusted substantially over this period.

The SIPP includes detailed information on individual and family demographic characteristics as well as the use of government transfer programs. The SIPP is a national survey that oversamples low-income households, but is nationally representative when weighted by survey weights. The respondents are interviewed every four months and asked about monthly activities during the prior four months. These 4-month interview periods are called waves. The panels vary in length from 32 to 48 months and vary in size from roughly 20,000 to 40,000 households. This large number of individuals gives us a sizable number of transitions even though the panels are fairly short.

The units of analysis in our empirical analysis are spells of female headship and non-headship and spells of welfare receipt and non-receipt. We define a female head of family as a woman who is unmarried and living with related children aged 17 or less. Our definition of female heads excludes married women with absent spouses but includes women in cohabiting relationships. Both of these cases are ambiguous because the contributions of the partners are not clear. We also include mothers who are heads of subfamilies. A narrower focus on female household heads would introduce an additional residential location component into the analysis. We define an indicator for headship and then compute spells of headship and non-headship based on the entire monthly sequence of headship indicators during course of the panel. Spells of headship and non-headship are mutually exclusive but possibly alternately repeating. That is, an individual woman could be observed to make multiple transitions into and out of headship and could contribute several spells to our analysis.

We include only spells of headship or non-headship that begin during the panel, that is, those that are uncensored or only right- censored. Although the SIPP contains retrospective information that can be used to construct a woman's entire headship history, it lacks retrospective information on many of the time-varying explanatory variables, particularly residence of children. It was not computationally feasible to account for these missing data. While excluding left censored spells leads to considerable sample loss, it correctly produces a sample of new spells to which our results apply. Several previous dynamic studies, including those by Blank & Ruggles (1996), Fitzgerald (1995) and Gittleman (2001), have also excluded left-censored spells.

We construct spells of welfare participation and non-participation in a similar way based on the monthly data. We define a woman as a participant if she receives AFDC or TANF income as the unmarried head of a family or subfamily unit. Conceivably, welfare spells could be defined independently of female headship. However, eligibility is conditioned on parenthood; so, the analysis would need to distinguish between different types of "non-heads." Also, there are not enough participating married families in the SIPP to support an empirical analysis. Because of this restriction, we do not examine participation in the Unemployed Parent programs of AFDC and TANF. Spells of welfare receipt or non-receipt that are on-going at the start of a headship spell are artificially left-censored at that point. Similarly, spells of welfare receipt, or non-receipt that are observed to continue after a woman exits headship are artificially right-censored. Within a given headship spell, there may be a single spell of welfare receipt or non- receipt or multiple spells on and off welfare.

In addition to the demographic and welfare information used to construct spell histories, the SIPP provides personal information such as age, race, and education, and whether the women lives in a metropolitan area.

The SIPP also provides information on geographic residence. We need residence information to assign values for welfare policy and labor and marriage market conditions, all of which vary by location and time. The public use version of the SIPP does not release county of residence nor does it fully report state of residence or MSA in order to preserve respondent confidentiality. To separate out the impact of welfare rule changes and labor and marriage market changes, we desire county of residence data so that we can use county variation to isolate labor market and marriage market effects. By special arrangement, we obtained permission to use the internal/confidential versions of the census files that reported county and state of residence. (The work was done at the US Census Bureau Boston Research Data Center and the Center for Economic Studies in Washington, DC. The results have been screened to insure that no confidential data are revealed and approved for release.) This permitted us to match in detailed contextual data.

Table 1 provides information about the characteristics of individuals in the top panel (a) followed by characteristics of spells in the bottom panel (b). We started with 88,419 women in the SIPP who were ever observed to be in the age range 15–54 and who ever reported any information on their headship or welfare histories. Of these, we could construct continuous, though possibly censored, spells with the necessary explanatory variables for 60,155 women. From this group, there were 12,685 women with at least one non-left-censored headship or non- headship spell. Thus, we lose about four-fifths of the available sample in the SIPP by dropping left-censored spells.

Our final sample of 12,865 women experienced 3,643 headship spells, 10,511 non-headship spells, 923 welfare participation spells, and 3,373 non-participation spells. Most of these spells are right censored. Previous estimates by Bumpass & Raley (1995) and Moffitt and Rendall (1995) indicated that the average length of headship spells was 3–4 years for whites and roughly 12 years for blacks; the high degree of right censoring in the SIPP is consistent with this. The non-headship spells in our sample are by younger individuals with less education because many of those spells begin with a woman aged 15. Welfare participants tend to be younger and less educated than non-participants. The second panel also displays time-varying characteristics of spells. In principle, the periodicity of the SIPP allows the time-varying characteristics to be updated every month. However, updating this information requires one record per spell; monthly updating would have produced roughly 300,000 records. To reduce the number of records with time-varying data, the analysis only updated these characteristics in the fourth month of every wave; this produced just under 80,000 time-varying, within-spell records. The numbers listed at the bottom of the table show the total number of time-varying records for each spell type.

Table 1. Means of the analysis variables

Variable				
Fixed individual characteristics				
Black	0.18			
Hispanic	0.13			
Number of individuals	12685			
	Female headship	Non-headship	Welfare participation	Welfare non-particip.
Spell characteristics				
Characteristics of spells				
Spell length	13.6	18.6	10.2	11.9
Proportion right censored	0.79	0.93	0.65	0.89
Age at start of spell	29.6	21.5	25.8	29.9
Education at start of spell	12.0	9.8	11.3	12.1
Number of spells	3643	10511	923	3373
Time-varying characteristics of spells				
Welfare participation	0.19	–	–	–
Maximum welfare benefit	351	355	379	345
Adopted any waiver	0.50	0.51	0.45	0.52
Term limit	0.14	0.15	0.08	0.16
Family cap	0.25	0.24	0.19	0.27
Teen coresidence requirement	0.23	0.23	0.18	0.24
AFDC-UP change	0.31	0.31	0.30	0.31
Work requirement	0.14	0.12	0.17	0.13
Earnings disregard change	0.29	0.28	0.29	0.28
JOBS change	0.24	0.25	0.15	0.26
Implemented TANF	0.38	0.38	0.27	0.41
Maximum EITC benefit	2530	2523	2375	2567
Log real wage in county	1.52	1.11	1.39	1.56
Employment probability in county	0.71	0.52	0.65	0.72
Sex ratio (male/female)	1.01	1.02	1.00	1.01
Metropolitan residence	0.81	0.79	0.82	0.81
Number of observations	13833	52032	2666	11167

Note: Figures calculated from the 1990, 92, 93 and 96 panels of the SIPP. Figures represent unweighted means unless otherwise indicated. Welfare participation and non-participation are measured only during headship spells.

Welfare policy parameters
The decision to receive welfare will depend on the level of benefits as well as other rules that affect eligibility. These will vary by state and over time. For benefits, we use the maximum benefit available for a family of three (Committee on Ways and Means, various years), deflated to 1992 dollars using the CPI-U. We choose a measure that does not vary by family size in order to avoid potential endogeneity of benefits based on fertility.

The remaining welfare policy parameters are indicators of specific rules. States experimented with many rule changes using waivers of federal policy up through 1996 when the PRWORA was passed. These waivers were adopted by different states at different times allowing us to identify their effects. TANF was also implemented at different times in different states, although within a narrow 14-month time window. We use information on waiver adoption and TANF adoption primarily from the US Department of Health and Human Services (DHHS) (1997) and Crouse (1999). The DHHS formed the waivers into main groups and determined when these were adopted statewide. Our measures include whether a state adopted any major waiver or whether it adopted specific waivers for a total lifetime limit on benefits (a termination limit), a reduced time limit before work was required (work time limit), reduced benefits for children who were conceived while the mother was on welfare (family cap), increased sanctions for failure to participate in the JOBS program or a reduction in the age of the youngest child for which the mother was required to participate in JOBS (JOBS sanctions), and more generous earned income disregards (earnings disregard).

We also used information on whether the state had relaxed rules for eligibility for the AFDC-UP program for married couples and whether a state had adopted a rule requiring teenage mothers to coreside with parents in order to receive benefits (the information on teenage coresidence requirements come from the Urban Institute's Welfare Rules Database). Finally, we defined an indicator for whether the state had implemented TANF. All of these indicators are time varying, with a value of zero prior to adoption and one thereafter based on the month and year of adoption.

Table 1 shows that a sizable amount of our observed spell time occurs after the adoption of some type of welfare waiver. In other work (Fitzgerald & Ribar 2004), we experimented with other variations on dating the waivers such as using implementation rather than adoption dates and using lags. Our overall results did not change substantially.

Besides welfare policy, we also include a variable that measures the generosity of the EITC. The EITC is a transfer that subsidizes earnings for low-wage workers and thus, interacts indirectly with welfare. The EITC was expanded substantially in the 1990s; several states have also adopted similar credits into their own income tax codes. The changes at the federal and state levels produce time series and cross-section variation in the value of the credit. We include a variable that measures the maximum combined federal and state credit for a family with two or more children, in 1992 dollars. The state EITC information was compiled by Nick Johnson at the Center on Budget and Policy Priorities.

Local labor and marriage markets
To measure job prospects, we impute county-level measures of skill- specific wages and employment probabilities by extending the work of Ribar (2003). In his work, Ribar constructed such measures for all counties from 1989 to 1997. He combined data from the Sample Edited Detail File (SEDF) of the 1990 Decennial Census and the 1990–1998 Annual Demographic files of the Current Population Survey (CPS) together with industry wage and employment information from the Regional Economic Information System (REIS). In order to identify county of residence and work, he used the internal/confidential versions of the SEDF and CPS by special arrangement with the Census Bureau. He estimated wages and probabilities of employment based on CPS and SEDF data on personal characteristics from those files as well as local employment and earnings measures from REIS. The selection-corrected wage regressions included county fixed effects and calendar time effects. We use these coefficients together with updated information from the REIS to impute wages and employment probabilities for women based on their county, education, age, and race over the period 1989–2000. We deflate earnings by the CPI-U. Table 1 shows the mean values. Predicted wages and

employment are smaller for the non-headship samples because of the lower average age of persons in those spells.

Since demographic decisions would be expected to depend on spouse availability, we construct a coarse measure of marriage market conditions. We use the ratio of men to women aged 15–39 in a county. Separate ratios were constructed for blacks, Hispanics and non-blacks using annual data from the 1990 decennial census and the intercensal county population estimates. Small samples in some counties led to lopsided numbers so we trim ratios that exceeded 5 or were less than 0.2 to those values.

Econometric specification

The study estimates hazard models of transitions from and into female headship and transitions from and into welfare participation. The transitions from female headship are specified to depend on welfare participation. The study applies Lillard's (1993) simultaneous hazards procedure to address problems of unobserved heterogeneity in all of the transition models and to account for the endogeneity of welfare participation in the headship model. The econometric specification is discussed in more detail below.

To examine the determinants of the timing of exits from female headship, the study estimates a log hazard model

$$\ln h_{\mathrm{H}}(t) = A_{\mathrm{H}}{}' T_{\mathrm{H}}(t) + \gamma P(t) + B_{\mathrm{H}}{}' X(t) + \eta. \qquad (1)$$

The hazard, $h_H(t)$, represents the probability of exiting female headship at month t conditional on having remained a head until at least t. In Equation (1), $T_H(t)$ represents a vector of duration parameters; P(t) is a time-varying indicator for welfare participation; X*(t)* is a vector of other observed and possibly time-varying covariates; η is an unobserved, person-specific variable, and A_H, γ, and B_H are coefficients. The first term on the right hand side of Equation (1), $A_H' \, T_H(t)$ is specified to be a linear spline in the spell duration. With this assumption, the hazard function has a piece-wise Gompertz specification.

The presence of unobserved heterogeneity in the hazard function is a substantial complication. Failure to account for such heterogeneity can lead to biased estimates of the coefficients (for instance, spurious indications of negative duration dependence). Following Lillard (1993), the study assumes that η is normally distributed with mean 0 and variance σ_η^2 and uses a maximum likelihood procedure that accounts for the distribution of headship spells under this assumption. The procedure is similar to the one developed by Butler and Moffitt (1982) for random-effect panel probit models in that it specifies the hazard function conditional on η and then integrates over the distribution and possible values of η.

Another complication is the endogenity of welfare participation. This problem is addressed by estimating models of headship and welfare participation jointly and allowing the unobserved determinants of these outcomes to be correlated.

Along with the model for exits from female headship, the study also estimates a model of the timing of entry into headship (exits from non-headship). The log hazard for this outcome is specified as

$$\ln h_{\mathrm{NH}}(t) = A_{\mathrm{NH}}{}' T_{\mathrm{NH}}(t) + B_{\mathrm{NH}}{}' X(t) + \lambda_{\mathrm{NH}}\eta \qquad (2)$$

where $T_{NH}(t)$ is a vector of duration parameters, *X(t)* and η are defined as before, and A_{NH}, B_{NH}, and λ_{NH} are coefficients. As with Equation (1), the log hazard for a spell of non-headship is specified as a piece-wise Gompertz distribution. The analysis allows for multiple, alternating spells of both headship and non-headship.

As Equations (1) and (2) indicate, a single unobserved factor is the source of unobserved heterogeneity in the hazard models for headship and non-headship. The coefficient λ_{NH} in Equation (2) relaxes the distribution somewhat. Without the coefficient (i.e., with $\lambda_{NH} = 1$), the sources of unobserved heterogeneity in the headship and non-headship models would be restricted to have the same variances and be perfectly, positively correlated. With the coefficient, the sources of unobserved heterogeneity in the two models can have different variances and be either perfectly positively or perfectly negatively correlated. While the single factor assumption clearly restricts the correlation between the sources of heterogeneity, it is adopted for reasons of tractability.

The log hazard functions for spells of welfare participation and non-participation are specified as

$$\ln h_{\mathrm{W}}(t) = A_{\mathrm{W}}' T_{\mathrm{W}}(t) + \Psi_{\mathrm{W}}' Z(t) + \mu \tag{3}$$

$$\ln h_{\mathrm{NW}}(t) = A_{\mathrm{NW}}' T_{\mathrm{NW}}(t) + \Psi_{\mathrm{NW}}' Z(t) + \lambda_{\mathrm{NW}} \mu \tag{4}$$

where $T_W(t)$ and TNW(t) are vectors of duration parameters, Z(t) is a vector of observed covariates, it is an unobserved, person-specific variable, and A_W, A_{NW}, Ψ_W, Ψ_{NM}, and λ_{NH} are coefficients, The unobserved variable μ is assumed to be normally distributed with mean 0 and variance σ_η^2. It is also assumed to be correlated with η (correlation coefficient ρ). The hazard equations for welfare participation and non-participation are only estimated during spells of headship; thus, they are conditioned on being a female head.

The four log hazard models are estimated jointly as a single system using the aML software package. The aML package employs Gaussian quadrature – a numerical approximation procedure – to evaluate the integrals over the two sources of unobserved heterogeneity. This study reports estimates from models that used eight quadrature points in each dimension, or 64 points total. Initial tests revealed that there were no noticeable differences in results between models that used six and eight points in each dimension.

Estimation results

Each of the models for female headship and welfare participation includes a piecewise linear specification for a baseline hazard. Preliminary models were estimated to determine the elements that would be included in $T_H(t)$, $T_{NH}(t)$, $T_W(t)$ and $T_{NW}(t)$–that is, to find the locations of the knots, or connections between segments, in the linear spline functions. To keep this initial specification search simple, the study restricted the elements of $T_H(t)$ and $T_{NH}(t)$ to be the same and restricted the elements of $T_W(t)$ and $T_{NW}(t)$ to be the same. Estimates from models with completely general duration patterns (dummy variables for each possible spell length) but no other controls guided the initial parameterizations of the piecewise linear baseline hazards. The study then added and deleted segments, checking to see whether these adjustments led to changes in the fit of the baseline models. The final baseline hazards for the headship and non-headship models were specified to have six segments corresponding to 0–3, 4–6, 7–9, 10–12, 13–30 and 31–48 months. The baseline hazards for the welfare participation and non-participation models were specified to have three segments corresponding to 0–3, 4–6 and 7–48 months. The specific elements of $T_H(t)$ and $T_{NH}(t)$ are

$$\begin{aligned}
&T_{0\text{–}3}(t) = \min(t, 3), \quad T_{4\text{–}6}(t) = \max[0, \min(t-3, 3)], \\
&T_{7\text{–}9}(t) = \max[0, \min(t-6, 3)], \quad T_{10\text{–}12}(t) = \max[0, \min(t-9, 3)], \\
&T_{13\text{–}30}(t) = \max[0, \min(t-12, 18)], \quad T_{31\text{–}48}(t) = \max(0, t-30).
\end{aligned}$$

The specific elements of $T_W(t)$ and $T_{NW}(t)$ are $T_{0\text{–}3}(t)$, T4–6(t), and $T_{7\text{–}48}(t)$ = max (0, t-6).

A similar procedure was employed to introduce a piecewise linear time trend into the models. The calendar time trend accounts for changes in national policies and socioeconomic conditions as well as differences across panels of the SIPP. The models for female headship and non-headship allow for different trends over the periods 1989–1991, 1992–1997 and 1998–2000 while the models for welfare participation and non-participation allow for different trends over the periods 1989– 1990, 1991–1998 and 1999–2000. The underlying variables for the trend segments are expressed in terms of calendar months since the end of 1988.

Table 2 reports coefficients for the welfare participation, welfare policy and local economic variables for three specifications of the system of transition models. The specifications differ in their controls for unobserved heterogeneity. The first column of Table 2 lists results from a specification that omits controls for unobserved heterogeneity. The second column lists results from a specification that includes controls for η and it but restricts these to be independent. The third column lists results from a specification that allows η and it to be correlated. For each specification, coefficients from the female headship hazard model are reported first; coefficients from the non-headship hazard model are reported second; coefficients from the welfare participation hazard model are reported third, and coefficients from the non-participation model are reported last. For brevity, Table 2 only reports a subset of coefficients from each model. In addition to the listed variables, the hazard models also include controls for race, ethnicity, age, education, metropolitan residence and the local sex ratio. Complete results for the specification reported in the third column of Table 2 are given in Appendix A. Complete results for the other specifications are available from the authors.

Estimation reveals that the controls for unobserved heterogeneity are statistically significant. In particular, the standard deviation for η in the headship model and the factor loading on η in the non-headship model are each individually distinguishable from zero (the factor loading is not statistically different from one, however). The corresponding parameters for μ are jointly but not individually significant. The positive factor loading in the female headship equation indicates that those prone to short spells of headship are also prone to short spells of non-headship. Thus, η appears to be associated with family instability generally. A similar interpretation applies to the coefficient in the welfare model, though this coefficient is insignificant. In the third specification, the correlation coefficient ρ is significantly negative. This indicates that characteristics that contribute to instability in living arrangements are associated with longer and more stable welfare program arrangements. Because specification tests reject the restrictions in the first two specifications, the discussion of empirical findings will focus on the coefficients from the third (least restrictive) specification. We note, however,

Table 2. Selected coefficients from Hazard models

	No controls for unobserved heterogeneity		Uncorrelated controls for unobserved heterogeneity		Correlated controls for unobserved heterogeneity	
Hazard for exiting headship						
Welfare participation	−0.48***	(0.12)	−0.50***	(0.12)	−0.43***	(0.13)
Adopted any waiver	0.08	(0.11)	0.08	(0.11)	0.08	(0.11)
Max. welfare benefits (/100)	−0.01	(0.03)	−0.01	(0.03)	−0.01	(0.03)
Max. EITC benefit (/1000)	−0.24*	(0.13)	−0.27**	(0.13)	−0.26**	(0.13)
Log real wage in county	−0.48**	(0.23)	−0.55**	(0.25)	−0.53**	(0.24)
Emp. prob. in county	0.59	(0.54)	0.84	(0.57)	0.81	(0.56)
σ_η	–		0.55***	(0.17)	0.42***	(0.13)
Hazard for entering or re-entering headship						
Adopted any waiver	−0.002	(0.11)	0.003	(0.11)	0.01	(0.12)
Max. welfare benefits (/100)	−0.02	(0.03)	−0.02	(0.04)	−0.03	(0.04)
Max. EITC benefit (/1000)	−0.38***	(0.14)	−0.41***	(0.14)	−0.42***	(0.14)
Log real wage in county	0.25	(0.30)	0.28	(0.32)	0.29	(0.33)
Emp. prob. in county	−0.40	(0.62)	−0.51	(0.65)	−0.56	(0.67)
λ_{NH}	–		1.19*	(0.64)	1.94**	(0.89)
Hazard for exiting welfare (conditional on headship)						
Adopted any waiver	−0.12	(0.16)	−0.13	(0.16)	−0.12	(0.16)
Max. welfare benefits (/100)	−0.09	(0.06)	−0.09	(0.06)	−0.09	(0.06)
Max. EITC benefit (/1000)	−0.61***	(0.20)	−0.61***	(0.21)	0.60***	(0.21)
Log real wage in county	−0.48	(0.46)	−0.47	(0.47)	−0.47	(0.47)
Emp. prob. in county	1.46*	(0.82)	1.46*	(0.83)	1.43*	(0.84)
σ_μ	–		0.15	(0.12)	0.11	(0.12)
Hazard for entering or re-entering welfare (conditional on headship)						
Adopted any waiver	−0.02	(0.16)	−0.03	(0.19)	−0.04	(0.18)
Max. welfare benefits (/100)	0.13***	(0.05)	0.14***	(0.06)	0.14**	(0.06)
Max. EITC benefit (/1000)	−0.71***	(0.18)	−0.79***	(0.20)	−0.78***	(0.20)
Log real wage in county	−0.91**	(0.39)	−1.00**	(0.47)	−0.95**	(0.47)
Emp. prob. in county	0.81	(0.71)	0.81	(0.87)	0.59	(0.88)
λ_{NW}	–		7.73	(6.89)	9.75	(11.1)
ρ	–		–		−0.38**	(0.19)
Log likelihood	−11229.84		−11207.10		−11205.04	

Note: Hazard models estimated using data from the 1990, 92, 93 and 96 panels of the SIPP. Models include splines for duration and calendar year effects and controls for race, ethnicity, age, education, metropolitan residence and sex ratio. Standard errors appear in parentheses.
* Significant at. 0.10 level.
** Significant at 0.05 level.
*** Significant at 0.01 level.

that the coefficients reported in Table 2 are not especially sensitive to the use of controls for unobserved heterogeneity.

Welfare participation is estimated to reduce the hazard of exiting female headship – that is, contribute to longer spells of headship. The estimated relationship is consistent with expectations and stronger than the preliminary results reported by Teitler et al. (2003). Estimates of the association between welfare participation and headship that account for correlations in the unobserved determinants in these outcomes are 10– 15% smaller than estimates that do not account for such correlations. Thus correcting for correlated heterogeneity makes a modest difference in the results.

Among the welfare policy variables, more generous welfare benefits are estimated to hasten unmarried mothers' participation in welfare. Benefits are also estimated to reduce exits from the welfare rolls, though the coefficient falls just below the threshold for statistical significance (two-tailed p value = 0.108). The coefficients for the welfare benefit variables are small and insignificant in the hazard models for headship and non-headship. Taken together, the estimates indicate that welfare benefits contribute indirectly to female headship by increasing welfare participation; however, there is no strong evidence of any additional direct association once participation is taken into account. None of the coefficients for the waiver variables is statistically different from zero. The weak results for waivers are consistent with our earlier findings for headship (Fitzgerald and Ribar, 2004) and welfare participation (Ribar 2004).

More generous benefits under the EITC are associated with longer spells of all four outcomes: headship, non-headship, participation and non-participation. Thus, the EITC appears to contribute to stability in both living and program arrangements. The finding that the EITC is associated with longer welfare spells is surprising but may reflect the subsidy allowing mothers to combine welfare and work careers. Previous research by Meyer & Rosenbaum (2001) indicated that the expansions in the EITC increased work but reduced welfare receipt, while research by Dickert-Conlin & Houser (1999) indicated that the subsidy reduced headship. Better economic opportunities in the respondent's county of residence in the form of higher average wages for women of the same age, race and schooling attainment significantly reduce the probability of exiting headship but also reduce the chances of entering welfare. The findings suggest that wage opportunities contribute to women's financial independence. The study's other measure of economic opportunities, the local skill-specific employment probability, is estimated to be positively associated with welfare exits.

The coefficients for the other observed variables (shown in Appendix A) either have the expected signs or are insignificant. In particular, women of African and Hispanic origin are generally estimated to have higher risks of headship and welfare participation than other women. The hazard for entry into female headship rises with age through age 18 then falls with age. The hazard for exiting headship increases with age, while the hazard for entering welfare falls with age. Higher levels of education help women avoid both headship and welfare participation. The hazards of exiting headship, non-headship and welfare participation increase with duration during the first three months of a spell. All four hazards generally decrease with duration after 4 months. The coefficients on the trend variables indicate that all four hazards were falling in the late 1990s.

Table 3 reports welfare participation, welfare policy and economic condition coefficients from three alternative specifications of the system of transition models. One issue that the study examines more carefully is whether the welfare policy variables have any independent effect on the duration of female headship once welfare participation is taken into account. The estimates from Table 2 indicate that the benefit level affects headship through welfare participation but that there are no additional independent effects of either the benefit level or waiver policies. The first column in Table 3 lists results from a specification that omits the welfare benefit and waiver variables from the headship equation. Other than these two exclusions, the specification includes all of the observed variables and statistical controls as the third specification from Table 2 (i.e., is nested within the previous specification). Thus, it can be used to test the joint significance of the policy variables in the headship model and examine the sensitivity of the welfare participation coefficient to their inclusion or exclusion. Comparisons across tables indicate that there is only a small change in the log likelihood function and no noticeable change in the coefficient for welfare participation, thus confirming our interpretation.

The second column in Table 3 lists results from a specification that adds an indicator for the implementation of TANF to each of the four hazard models. Most states had reformed their welfare programs through the Section 1115 waiver process by the end of 1996; however, as a result of the PRWORA all states were subsequently required to implement TANF programs. In some states, the TANF programs followed the general contours of the waiver provisions. In other states, TANF represented a substantial change in direction or the actual start of the reform process. Including indicators for both waiver adoption and TANF implementation provides a more complete description of

Table 3. Selected coefficients from Hazard models with different policy controls

	Controls for welfare participation only		Controls for TANF implementation		Control for detailed waiver provisions	
Hazard for exiting headship						
Welfare participation	−0.43***	(0.13)	−0.43***	(0.13)	−0.43***	(0.13)
Adopted any waiver	–		0.08	(0.11)	–	
Term limit	–		–		0.11	(0.16)
Family cap	–		–		0.01	(0.12)
Teen coresidence req.	–		–		0.11	(0.12)
AFDC-UP change	–		–		−0.02	(0.13)
Work requirement	–		–		−0.05	(0.17)
Earnings disregard change	–		–		−0.05	(0.12)
JOBS change	–		–		0.05	(0.13)
Implemented TANF	–		−0.07	(0.18)	−0.14	(0.19)
Max. welfare benefits (/100)	–		−0.01	(0.03)	0.01	(0.03)
Max. EITC benefit (/1000)	−0.27**	(0.13)	−0.29**	(0.14)	−0.32**	(0.15)
Log real wage in county	−0.53**	(0.22)	−0.53**	(0.24)	−0.54**	(0.25)
Emp. prob. in county	0.82	(0.54)	0.81	(0.56)	0.78	(0.58)
σ_η	0.42***	(0.12)	0.43***	(0.13)	0.45***	(0.13)
Hazard for entering or re-entering headship						
Adopted any waiver	0.01	(0.12)	0.01	(0.12)	–	
Term limit	–		–		0.11	(0.18)
Family cap	–		–		−0.004	(0.12)
Teen coresidence req.	–		–		0.24**	(0.12)
AFDC-UP change	–		–		−0.04	(0.15)
Work requirement	–		–		0.13	(0.19)
Earnings disregard change	–		–		−0.13	(0.12)
JOBS change	–		–		−0.07	(0.14)
Implemented TANF	–		0.03	(0.19)	−0.06	(0.20)
Max. welfare benefits (/100)	−0.03	(0.04)	−0.03	(0.04)	−0.02	(0.04)
Max. EITC benefit (/1000)	−0.42***	(0.14)	−0.41***	(0.15)	−0.46***	(0.16)
Log real wage in county	0.29	(0.33)	0.29	(0.33)	0.33	(0.33)
Emp. Prob. in county	−0.55	(0.67)	−0.56	(0.67)	−0.68	(0.68)
λ_{NH}	1.93**	(0.89)	1.90**	(0.89)	1.80**	(0.86)
Hazard for exiting welfare (conditional on headship)						
Adopted any waiver	−0.12	(0.16)	−0.12	(0.16)	–	
Term limit	–		–		−0.57*	(0.31)
Family cap	–		–		0.13	(0.18)
Teen coresidence req.	–		–		0.14	(0.19)
AFDC-UP change	–		–		0.15	(0.24)
Work requirement	–		–		−0.45	(0.30)
Earnings disregard change	–		–		−0.13	(0.21)
JOBS change	–		–		0.12	(0.21)
Implemented TANF	–		−0.17	(0.24)	−0.20	(0.26)

Table 3. Continued

	Controls for welfare participation only		Controls for TANF implementation		Control for detailed waiver provisions	
Max. welfare benefits (/100)	−0.09	(0.06)	−0.09	(0.06)	−0.08	(0.06)
Max. EITC benefit (/1000)	−0.60***	(0.21)	−0.66***	(0.23)	−0.82***	(0.24)
Log real wage in county	−0.47	(0.47)	−0.47	(0.47)	−0.57	(0.47)
Emp. prob. in county	1.43*	(0.84)	1.42*	(0.84)	1.62*	(0.85)
σ_μ	0.11	(0.12)	0.11	(0.12)	0.11	(0.13)
Hazard for entering or re-entering welfare (conditional on headship)						
Adopted any waiver	−0.04	(0.18)	−0.03	(0.19)	–	
Term limit	–		–		0.18	(0.33)
Family cap	–		–		0.37*	(0.21)
Teen coresidence req.	–		–		−0.001	(0.22)
AFDC-UP change	–		–		0.23	(0.25)
Work requirement	–		–		0.46	(0.30)
Earnings disregard change	–		–		−0.28	(0.23)
JOBS change	–		–		−0.60**	(0.25)
Implemented TANF	–		−0.36	(0.26)	−0.32	(0.28)
Max. welfare benefits (/100)	0.14**	(0.06)	0.14**	(0.06)	0.10*	(0.06)
Max. EITC benefit (/1000)	−0.78***	(0.20)	−0.88***	(0.23)	−0.81***	(0.23)
Log real wage in county	−0.94**	(0.47)	−0.96**	(0.47)	−0.83*	(0.49)
Emp. prob. in county	0.59	(0.88)	0.64	(0.88)	0.52	(0.90)
λ_{NW}	9.75	(11.1)	10.0	(11.4)	9.82	(11.4)
ρ	−0.38**	(0.19)	−0.36*	(0.19)	−0.34*	(0.19)
Log likelihood	−11205.35		−11203.63		−11185.48	

Note: Hazard models estimated using data from the 1990, 92, 93 and 96 panels of the SIPP. Models include splines for duration and calendar year effects and controls for race, ethnicity, age, education, metropolitan residence, and sex ratio. Standard errors appear in parentheses.
* Significant at 0.10 level.
** Significant at 0.05 level.
*** Significant at. 0.01 level.

reform efforts and allows for differences between waiver and TANF policies. The estimated coefficients for TANF, however, are all statistically weak, and a likelihood ratio test indicates that they are jointly insignificant. The strongest result appears in the hazard model for non- participation. The coefficient suggests that TANF may have slowed and reduced entry into welfare, but the p value is only 0.175.

The third specification replaces the single indicator for adopting any type of welfare waiver with seven separate indicators for different types of waivers. It is reasonable to expect that some types of waivers might have stronger or weaker effects, or possibly even differently signed effects, on headship and participation outcomes. The third specification allows for such effects but at the potential expense of a loss of statistical power if the policies are closely related or only implemented in a few locations. Estimation reveals that few of the individual waiver indicators are statistically significant (only four coefficients out of the 28 entered into the models). Of the coefficients that are significant, most have counter-intuitive signs. For instance, teen coresidence requirements are associated with faster entry into headship, and term limits are associated with longer spells on welfare. The results provide little support for the hypothesis that waiver provisions played a meaningful role in the stabilization of headship rates or in the decline in welfare participation.

Conclusion

This study draws individual-level data on spells of female headship, non-headship, welfare participation, non-participation from several panels of the SIPP. Through an arrangement with the U.S. Census Bureau, we use special versions of the SIPP that allow us to link these data with state-level indicators of welfare policies and county-level measures of economic and marriage opportunities. The study uses the combined data to estimate hazard models of the four spell outcomes. The estimation procedure accounts for correlated sources of unobserved heterogeneity in the determinants of spell lengths. The procedure also allows the study to consider welfare participation as an endogenous determinant of female headship spells.

Estimates from the hazard models indicate that welfare participation is significantly, negatively associated with the probability of leaving headship. This association is robust in terms of sign, magnitude and statistical significance to the use of controls for endogeneity. The finding is consistent with welfare participation directly contributing to longer spells of female headship. While the evidence regarding causality is stronger than that reported in some previous studies, it is not definitive because the study's statistical methodology only accounts for endogeneity that arises from unobserved, time-invariant characteristics of people and relies on relatively strong assumptions regarding the distribution of these unobserved characteristics.

The study finds that the chances that an unmarried mother will enroll in welfare increase with the level of benefits offered by her state of residence. There is also weak evidence that benefits encourage unmarried mothers to remain on welfare. More generous benefits are indirectly associated with longer spells of female headship through their association with welfare participation. The study does not find evidence that benefits have an additional, direct impact on headship, once the effect through welfare participation is taken into account. Other welfare policies, as measured by the adoption of program waivers and the implementation of TANF, are not strongly associated with female headship or welfare participation. Thus, aside from changes in benefits, it does not appear that reforms enacted during the 1990s contributed substantially to the stabilization of headship rates or the reduction in welfare caseloads.

A strength of this study is its use of skill-specific, county-level controls for wage and employment opportunities. Higher wages are associated with longer spells of female headship as well as longer spells off welfare. These results, along with like-signed estimates for EITC benefits, suggest that earnings contribute to women's economic independence – both from potential husbands and from the welfare system.

Some limitations of the study should also be kept in mind in interpreting the results. The biggest limitation is the short observational window available in the SIPP. In no instance could the study examine headship or welfare participation spells that lasted more than 4 years. Shorter time frames for some panels, attrition from the surveys, and the study's exclusion of initially on-going (left-censored) spells further limited the number of transitions that could be examined. While the breadth of coverage makes the SIPP a logical choice for examining welfare policies, surveys with a greater length of coverage, such as the PSID, should also be considered in future work. Other limitations of the study include the strong parametric assumptions in the hazard functions and the complexity of the estimation methods. Despite these limitations, we are confident that our study contributes to an emerging consensus that welfare reform has had no more than a modest effect on demographic outcomes.

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Note
1. Some studies, however, document that bad marriage situations may be worse for children than single parenthood (Amato et al. 1995; Hofferth & Anderson 2003; Jekielek 1998; Morrison & Coiro 1999). Deleire & Kalil (2002) argue that single- parent multigenerational families can be beneficial. Further,

Amato's (1993) review of the literature notes that many studies find insignificiant associations between family structure and child well being.

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Appendix A. Full results from preferred Hazard model

	Exiting Female headship		Entering or re-entering headship		Exiting welfare		Entering or re-entering welfare	
Linear spline for duration								
0–3 months	1.20***	(0.12)	1.21***	(0.15)	0.62***	(0.12)	−0.08	(0.08)
4–6 months	−0.42***	(0.06)	−0.22***	(0.07)	−0.23***	(0.07)	−0.22***	(0.07)
7–9 months	0.10	(0.07)	0.14*	(0.08)	-		–	
10–12 months	−0.13**	(0.07)	−0.24***	(0.07)	–		–	
13–30 months	0.01	(0.01)	0.02	(0.01)	–		–	
31–48 months	−0.13*	(0.08)	0.03	(0.03)	–		–	
7–48 months	–		–		−0.001	(0.01)	−0.04*	(0.02)
Linear spline for time trend								
1989–1991	−0.03**	(0.01)	0.02	(0.01)	–		–	
1992–1997	0.01	(0.005)	0.01**	(0.005)	–		–	
1998–2000	−0.02***	(0.01)	−0.02***	(0.01)	–		–	
1989–1990	–		–		0.001	(0.18)	0.24***	(0.09)
1991–1998	–		–		0.03***	(0.01)	0.02**	(0.01)
1999–2000	–		–		−0.08**	(0.03)	−0.08**	(0.03)
Other covariates and controls								
Welfare participation	−0.43***	(0.13)	-		-		-	
Adopted any waiver	0.08	(0.11)	0.01	(0.12)	−0.12	(0.16)	−0.04	(0.18)
Max. welfare ben. (/100)	−0.01	(0.03)	−0.03	(0.04)	−0.09	(0.06)	0.14**	(0.06)
Max. EITC benefit (/1000)	−0.26**	(0.13)	−0.42***	(0.14)	−0.60***	(0.21)	−0.78***	(0.20)
Log real wage in county	−0.53**	(0.24)	0.29	(0.33)	−0.47	(0.47)	−0.95**	(0.47)
Emp. prob. in county	0.81	(0.56)	−0.56	(0.67)	1.43*	(0.84)	0.59	(0.88)
Sex ratio in county	0.11	(0.37)	0.30	(0.32)	–		–	
Metropolitan residence	−0.06	(0.10)	0.10	(0.12)	−0.12	(0.17)	−0.01	(0.17)

Appendix A. Continued

	Exiting Female headship		Entering or re-entering headship		Exiting welfare		Entering or re-entering welfare	
Age at start of spell	0.04***	(0.01)	−0.05***	(0.01)	0.001	(0.01)	−0.05***	(0.01)
Age = 15	–		−3.16***	(0.20)	–		–	
Age = 16	–		−2.55***	(0.27)	–		–	
Age = 17	–		−0.17	(0.42)	–		–	
Age = 18	–		0.71*	(0.39)	–		–	
Education at start of spell	−0.03	(0.02)	−0.05*	(0.03)	0.02	(0.05)	−0.12**	(0.05)
Black	−0.48***	(0.11)	0.77***	(0.11)	−0.48***	(0.15)	0.40***	(0.15)
Hispanic	−0.30**	(0.12)	0.44***	(0.12)	−0.52***	(0.17)	−0.19	(0.18)
Intercept	−5.66***	(0.70)	−6.00***	(0.76)	−4.60	(4.25)	−6.05***	(2.11)
σ_η, σ_μ	0.42***	(0.13)	–		0.11	(0.12)	-	
$\lambda_{NH}, \lambda_{NW}$	–		1.94**	(0.89)	-		9.75	(11.1)
ρ			−0.38**		(0.19)			
Log likelihood					−11205.04			

Note: Hazard models estimated using data from the 1990, 92, 93 and 96 panels of the SIPP.
Standard errors appear in parentheses.
* Significant at 0.10 level.
** Significant at 0.05 level.
*** Significant at 0.01 level.